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U.S. COMMERCIAL BANKS AND THE EURODOLLAR MARKET*

This paper presents an econometric analysis of U.S. commercial bank borrowing in the Eurodollar market. Section I briefly reviews institutional and econometric considerations. Our data sources, definitions and transformations are described in Section II. In Section III we consider the issues of simultaneity-bias correction, of including a wide range of asset and liability returns as explanatory variables, and of the specification of the lag structure. Section IV suggests some economic implications of our results.

I. General Considerations in Specification and Estimation

The position of U.S. commercial banks in international financial markets differs from that of many non-American banks. While other banks are involved in transactions in assets and liabilities denominated in currencies other than their national ones, U.S. commercial banks seem to deal primarily in financial instruments denominated in dollars.[1] We focus our study on the

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[1] The exact position of U.S. banks is difficult to assess since data on their activities is inadequate. However, the series of liabilities of large city banks to their foreign branches is assumed by both institutionally (Einzig [1970]; Clendenning [1970]) and econometrically (Black [1971]; Marston [1974]) oriented writers to be primarily composed of Eurodollar borrowings by these banks. (See Section II for further details.)

Eurodollar market, an important source of international funds for U.S. commercial banks.

As has been emphasized by Leamer and Stern [1972] critical importance must be attached to an appropriate disaggregation of transactor and financial instruments. In turn, financial instruments should be distinguished as to maturity, country of origin and other characteristics. It is generally held that of all transactors, banks are the most susceptible of analysis using the portfolio approach we have adopted.

A second issue in disaggregation is the specification of the asset and liability menu of the transactor. In this study we emphasize the fairly rich menu facing banks. As a consequence, portfolio adding-up constraints suggest that all the rates on these alternative financial instruments appear in the equation for Eurodollar borrowing (Brainard and Tobin [1968]). To our knowledge this study presents the first attempt to discuss the importance of specifying a wide menu in the context of the international capital markets literature, and we present results which support this strategy.

A third issue in disaggregation involves the region of origin and currency of denomination of financial instruments. The difficult question of foreign exchange risk is closely associated with this aspect of disaggregation. In the context of our stylized representation in which U.S. commercial banks choose between domestic instruments and Eurodollar liabilities, foreign disaggregation need be taken no further than the relatively homogeneous Eurodollar market. Furthermore, the question of exchange risk is of secondary priority since these liabilities are largely dollar denominated.

The treatment of the exogenous component of the transactor's balance sheet--often termed the scale variable--is another important issue. Portfolio theory (Brainard and Tobin [1968]) imposes some minimal requirements on the treatment of this variable in specification. In the context of these restrictions, a popular functional form applied to our problem is

$$\frac{EDB}{W^*} = ar$$

where EDB is Eurodollar borrowing, W^* is the scale variable, r is a vector of interest rates and a is a coefficient vector. The ratio formulation assumes that the vector of interest rates determines the share of the portfolio that is held in a particular asset or liability.

A second aspect of the functional form is the specification of lag structure. One justification for a lagged response to interest rates by commercial banks is associated with the formation of expectations. Past values of interest rates may affect expectations of future rates. A second justification for past values of variables appearing in the equation is slow adjustment. Lock-in problems associated with past financial decisions and commitments may generate a lagged response. If liquidity is problematic for financial instruments other than Eurodollar liabilities such lock-in effects may spill-over, via adding-up constraints, into banks' demands for Eurodollars (Brainard and Tobin [1968]). Thus Eurodollar borrowing will itself be a function of the past values of interest rates on the basis of which past financial commitments were made.

Finally, in the presence of errors-in-variables, improved estimations may result from the inclusion of lagged values of the independent variables. A further discussion of this problem appears in Section III.

While these general considerations point to the inclusion of lagged values as explanatory variables, a specific formulation must still be adopted. One possible approach is to include the lagged value of the dependent variable as an explanatory variable. This partial adjustment approach unfortunately may have undesirable implications in a balance-sheet context (Brainard and Tobin [1968]). Furthermore, a partial adjustment formulation imposes identical responses of the dependent variable to changes in all the independent variables. Our approach has therefore been to adopt a polynomial distributed lag formulation which overcomes these difficulties.

A final estimation problem which we treat is simultaneous equations bias. The importance of U.S. commercial banks in the Eurodollar market suggests that their activity as a group may well affect the rates at which their borrowing takes place. A randomly high level of Eurodollar borrowing by U.S. banks may raise the rate at which such borrowing takes place. As a result of this positive correlation between the random component of the dependent variable and the Eurodollar rate as an explanatory variable the demand elasticity estimated by non-simultaneous equations techniques will be biased toward zero.

Despite these _a priori_ arguments, the predominent attitude in the empirical literature on international capital flows has been to ignore the simultaneity problem. Indeed, some authors such as Leamer and Stern [1972] have explicitly considered a treatment of this problem to be of decidedly secondary importance. Nonetheless, two important studies, Black [1971] and Marston [1974], treat simultaneity by reduced form estimation. An unfortunate aspect of this approach is that it requires a correct specification of the behavioral relations for participants in the Eurodollar market other than U.S. commercial banks. Since many of the non-American transactors are subject to complicated regimes of financial regulation, a correct specification of the non-American side of the market is exceedingly difficult (OECD [1972, 1973 a and b]). It is unlikely that two or three European interest rates can adequately portray the factors influencing non-American participants in the market. Failure to specify correctly the non-American side of the market results in <u>inconsistent</u> estimates when a reduced form technique is used.

Further, these two studies, as well as Hewson and Sakakibara [1974] another simultaneous technique study, consider European interest rates as exogenous. If, on the other hand, European rates are assumed dependent on Eurodollar flows these studies have only partially corrected for simultaneity. Of course, reduced-form estimation has a number of advantages, especially in providing a simultaneously-determined system. However, our strategy of disaggregation makes infeasible a fully-specified reduced-form model, and we proceed with a structural equation technique which will allow us to consider the importance of a simultaneity bias correction.

II. Definition of Variables and Period of Estimation

This study examines the average monthly liabilities of U.S. large city commercial banks (LCB's) to their foreign branches between November 1964 and December 1971. The explanatory variables are listed in Table 1. Our explanatory variables include a vector of four U.S. interest rates,[2] a scale variable, variables representing unanticipated changes in deposits and loans, the Regulation M reserve requirement, and monthly dummies.

Items on the banks' balance sheets which are exogenous to the banking sector because effective interest rate ceilings prevail, such as demand and time deposits, are incorporated in the scale variable. To isolate trends from short-run fluctuations of this variable, it is entered as an arithmetically-weighted average of the current and past fifteen months' observations.

Unanticipated short-run fluctuations in the banks' exogenous liabilities have potentially quite different effects on Eurodollar borrowing than long-run changes in this variable. Similarly banks may be faced with short-run demand for loans which they satisfy without simultaneously adjusting their lending rates. The Eurodollar market serves as a potential source of funds for such loans. We use an arithmetically-weighted average of current and past values of exogenous liabilities and gross adjusted loans to represent trends in these variables. Explanatory variables representing short-term variations in these quantities are defined as the differences between their contemporaneous and trend values.

One difficulty in estimating Eurodollar borrowing by banks in this period is the Regulation Q ceiling on interest paid on certificates of deposit. For

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[2]Following (Brainard and Smith [1974]) all interest rates are entered as inverses.

31 of the original 86 months in the estimation sample the rate in the secondary market for CD's exceeds the ceiling.[3] During these months in which the ceiling is effective CD's are analogous, from the bank's point of view, to time and demand deposits in that the quantities they receive are less than their desired level at the ceiling rate. CD's are then appropriately classified as exogenous liabilities and should therefore be included in the scale variable. Furthermore, the CD interest rate should not be considered an explanatory variable. For these reasons we conclude that an effective ceiling on the CD rate creates a qualitatively different banking environment requiring separate estimation. Hence we exclude from our estimation period those 31 months in which the market CD rate exceeds the ceiling. Alternatively a general simultaneous equations model of the U.S. financial sector could be developed to incorporate Eurodollar borrowing behavior under either regime. We consider such an effort to be outside the scope of this study.

Regulation M, which imposed a reserve requirement on liabilities of U.S. commercial banks to their foreign branches above a certain level presents a further estimation problem. The reserve-free level for each bank depended positively on its Eurodollar borrowing in previous periods, suggesting that the shadow cost of Eurodollars has occasionally been less than the interest cost adjusted for the reserve requirement. To capture this effect the reserve requirement ratio imposed by Regulation M is included as a separate explanatory variable.[4]

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[3] Comparison showed the CD ceiling to be effective for: July through December 1966, December 1967, and July 1968 through June 1970.

[4] See the Federal Reserve Bulletin for a discussion of the rate complicated reserve requirements imposed by Regulation M.

III. Sensitivity of the Analysis to Alternative Estimation Strategies

We estimate an equation of the form

$$(3.1) \qquad \frac{EDB}{W^*}_t = a + \sum_{i\epsilon I} \sum_{j=0}^{15} b_{ij}(1/i)_{t-j} + cRR^{ed}_t + d\left(\frac{W - W^*}{W^*}\right)_t + e\left(\frac{L - L^*}{W^*}\right)_t + \sum_{j=1}^{11} fjDj_t + u_t$$

where $I = (r^{ed}, r^{ff}, r^{cd}, r^{gs}, r^{pr})$. Variable definitions appear in Table 1. We impose an Almon second-order polynomial distributed lag structure on the b_{ij}'s . The coefficient of the contemporaneous Eurodollar rate is treated separately to allow correction for potential simultaneity bias. Thus

$$b_{ij} = a_{0i} + a_{1i}j + a_{2i}j^2$$

for the U.S. rates while

$$b_{ij} = a_{0i} + a_{1i}(j-1) + a_{2i}(j-1)^2$$

for $i = r^{ed}$. We constrain b_{i15} , the coefficient of the farthest observation, to equal zero, so that

$$a_0 = -ka_1 - k^2a_2$$

where $k = 15$ for the U.S. rates and $k = 14$ for r^{ed} . With this lag scheme incorporated our equation may be rewritten as

$$(3.1') \qquad \frac{EDB}{W^*}_t = a + b_{edt}(1/r)^{ed}_t + \sum_{i\epsilon I} [b'_iX1^i_t + b''_iX2^i_t] + cRR^{ed}_t + d\left(\frac{W - W^*}{W^*}\right)_t + e\left(\frac{L - L^*}{W^*}\right)_t + \sum_{j=1}^{11} fjDj_t + u_t$$

where

$$X1_t^i = \sum_{j=1}^{14} (j - 15)(1/r)_{t-j}^{ed}$$

$$X2_t^i = \sum_{j=1}^{14} (j^2 - 197)(1/r)_{t-j}^{ed}$$

for $i = r^{ed}$ and

$$X1_t^i = \sum_{j=0}^{14} (j - 15)(1/r)_{t-j}^{i}$$

$$X2_t^i = \sum_{j=0}^{14} (j^2 - 225)(1/r)_{t-j}^{i}$$

otherwise.

Equation (3.1') was estimated by ordinary least squares for the period January 1964 through December 1971. Months in which the Regulation Q ceiling on the CD rate were effective are omitted from the estimation sample. Table 2 reports the point estimates of the coefficients and summary statistics of the regression.

To assess the sensitivity of our results to alternative estimation approaches we present departures in four directions from the procedure described above. Specifically we consider the effect of (1) correcting for simultaneity bias in the contemporaneous Eurodollar rate, (2) reducing the length of the lag structure from 15 to 6 months, (3) including a smaller number of domestic rates as explanatory variables, and (4) including months in which the Regulation Q CD rate ceiling is effective in the estimation period.

1. To correct for potential simultaneity bias in the estimation of equation (3.1') we apply a variant of two-stage-least-squares estimation.

Brundy and Jorgenson [1974] have shown that consistent but possibly inefficient estimators can be obtained by using a subset of all exogenous variables in the model to adjust endogenous right-hand side variables. Our instrument list includes all exogenous variables included in equation (3.1'). In addition we employ lagged values of foreign industrial production indices to reflect the general level of economic activity in the major non-U.S. participants in the Eurodollar market. Such activity will affect in various ways the supply of Eurodollars. The absence of autocorrelation in our estimated equation justifies the use of lagged endogenous variables as instruments.

The long-run elasticities yielded by ordinary least squares and by the instrumental variables estimation technique may be compared in Table 3. The results seem insensitive to our attempt to correct for simultaneity bias, supporting the view of Leamer and Stern [1972].

2. We consider the effect of shortening the length of the lag to 6 months and 2 months, maintaining an Almon polynomial distributed lag structure with the far zero constraint. Estimation with either specification yields evidence of a severe autocorrelation problem. With the 6-month (2-month) lag the Durbin-Watson statistic before autocorrelation adjustment is 1.061 (0.37) while the Cochrane-Orcutt estimate of the first-order autocorrelation coefficient is .91 (.92) with a t-statistic of 15.4 (16.3). These results may be symptomatic of misspecification or an inadequate treatment of measurement error.[5] By way of contrast, equation (3.1') yields a Durbin-Watson statistic of 2.01. The corresponding estimate of first-order autocorrelation is insignificantly

[5]Grether and Maddala [1973, pp. 256-257] have shown in a simple model that, even if there is no serial correlation in the true error of the equation, serial correlation of either the true value of the independent variable or its measurement error will produce calculated residuals which are autocorrelated.

different from zero.

Furthermore, the R^2 statistic with the 6-month (2-month) lag specification is .9666 (.8288) while with the 15-month lag it is .9947. In both cases extending the lag backward increases the R^2 statistic dramatically. Note that because we employ a second order polynomial in all cases the number of explanatory variables is the same in each case. Thus the R^2 are completely comparable.[6] We conclude from our examination of Durbin-Watson statistics, R^2's and econometric theory that a long lag structure is appropriate.

We also note a tendency for the estimated long-run elasticities to decline when the lag is shortened. Especially dramatic is the comparison of the more traditional 2-month lag with the 15-month specification. In this case the estimate of the own rate elasticity drops by a factor of 5 when the lag is reduced.

We refrain, however, from making inferences about the true form of the lag structure. Grether and Maddala [1973, p. 258] have shown in a simple model, that errors-in-variables may exaggerate the length of the lag.[7] Nonetheless, under conditions likely to be satisfied in our model, inclusion of lagged independent variables will improve the estimate of the total (i.e. long-run) effect.[8]

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[6] In each case the dependent variable is indexed over the period described in Section II. Hence, the degrees of freedom are the same for each regression.

[7] For this reason a detailed discussion of the estimated lag structure is omitted. The profiles of the estimated lags may be derived from the coefficients presented in Table 2. We do point out, however, that 50 per cent of the total effect is obtained by 6, 5, 4, 3 and 10 months for r^{ed}, r^{ff}, r^{cd}, r^{gs} and r^{pr} respectively.

[8] For a further discussion of problems in interpreting summary statistics of distributed lag estimates, see Sims [1971, 1972]. Sims has shown that when continuous time processes must be estimated with data observed discretely, inclusion of observations of explanatory variables outside the range indicated by the underlying continuous time process may yield improved estimates of the total effect.

Another aspect of our lag specification is the far zero constraint on the interest rates. A joint test on the five constraints failed to reject the null hypothesis that relaxing these constraints has no significant effect. We conclude that these restrictions are not contraindicated by the data.

3. Since the inclusion of a large number of interest rates is a procedure not generally followed in the literature, we consider it worthwhile to examine the effect of reducing the list of interest rates on the point estimate of the own-rate elasticity. Table 4 reports the estimate of the own-rate elasticity when each domestic rate is dropped individually and when the liability and asset rates are dropped in pairs. Omitting the prime rate alone lowers the point estimate of the own-rate elasticity considerably. If both asset or both liability rates are omitted the decline is even more dramatic. We interpret these results as suggestive of missing variables bias when only domestic asset or only domestic liability rates appear as explanatory variables.[9]

4. A Chow test rejects at the one per cent confidence level the null hypothesis that the periods in which the Regulation Q CD rate ceiling was effective are homogeneous with our estimation period. We conclude that our omission of these observations from the sample is justified.

In summary we conclude that equation (3.1'), embodying a long lag structure and a comprehensive interest rate list, represents an appropriate speci-

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[9]The expected value of the missing variable bias is

$$(X_1'X_1)^{-1} \quad (X_1'X_2)a_2$$

where X_1 is the observation matrix of included variables, X_2 are improperly excluded variables and a_2 is the vector of the true coefficients of the excluded variables. In our context positive coefficients of the regression of excluded on included variables will bias the own-rate elasticity toward zero, the result we observe.

fication of U.S. Commercial Bank borrowing in the Eurodollar market. We also conclude that correction for simultaneity bias in the estimated coefficient of the contemporaneous Eurodollar rate is not indicated, that the zero constraint on the coefficient of the far month is appropriate, and that months in which the Regulation Q ceiling is effective should not be pooled with our sample.

IV. Economic Interpretation of the Estimation Results

This section presents an analysis of the significance levels of our estimates and suggests a number of inferences that may be made about the behavior of U.S. commercial banks in the Eurodollar market.

We tested the null hypothesis that the long-run interest rate coefficients are zero. The joint test on all five interest rates yields an F ratio which allows rejection of the null hypothesis at the 99 per cent level. When tested singly each long-run interest rate coefficient is significant at the 95 per cent confidence level except the CD rate coefficient. All but the CD and 3 to 5 year government security rate coefficients are significant at the 99 per cent level as well. These results are particularly remarkable in the context of financial market estimation where the presumption is that multicollinearity leads to standard errors so large as to prohibit interest rate coefficients from being individually significant.

We offer the following observations on results:

1. The estimated elasticities indicate that large changes in the position of U.S. commercial banks in the Eurodollar market accompany relatively small changes in interest rates. Table 5 presents estimates of the long-run elasticities. The numbers in parentheses are slopes, or the dollar change in LCB Eurodollar liabilities resulting from a one-hundred basis point change in the indicated interest rates, other rates held constant.

2. Except for the insignificant CD rate coefficient, the point estimates of the interest rate coefficients are compatible with the assumption frequently made in portfolio analysis that financial instruments are gross substitutes.[10] The point estimate of the own-rate elasticity is negative while those for all cross elasticities except for that of the CD rate are positive and smaller in absolute value at the mean.

[10] See, for instance, Brainard and Tobin [1968].

3. A one per cent (one-hundred basis point) rise in all domestic rates accompanied by a one per cent (one-hundred basis point) rise in the Eurodollar rate from mean values increases Eurodollar liabilities by 4.1 per cent ($4.5 billion). A positive effect on borrowing of an equal rise in both domestic and foreign interest rates is also reported by Branson and Willett [1972].

4. We find no evidence that Eurodollar liabilities react to changes in U.S. liability rates more than to changes in asset rates. The point elasticities indicate that an equal percentage increase in liability and decrease in asset rates will not increase Eurodollar indebtedness.

5. We attempt to estimate the impact effect of deviations in exogenous liabilities and loans from their trend levels by including $(W - W^*)/W^*$ and $(L - L^*)/W^*$ as explanatory variables. Unanticipated variation in exogenous liabilities does not seem to have served as an important reason for changes in Eurodollar liabilities. The coefficient is not significantly different from zero and is of the unanticipated sign. The result indicates that a $1.00 fall in net demand and time deposits is associated with a 2 cent <u>fall</u> in Eurodollar liabilities.

The Eurodollar market does seem to have served as an important source of funds for satisfying off-trend loan demand. The coefficient of $(L - L^*)/W^*$ is significantly positive. Its point estimate indicates that for each dollar of such unanticipated loans, 19 cents is borrowed from overseas branches. This result contrasts sharply with the <u>average</u> ratio of Eurodollar liabilities to loans in this period, which is only 2.5 cents per dollar.

6. The Regulation M reserve requirement has apparently had a negative overall effect on Eurodollar liabilities. The total effect may be divided into a negative cost-of-borrowing effect resulting from the increase in the adjusted Eurodollar rate, holding the unadjusted Eurodollar rate constant,

and a positive demand-for-reserves effect due to the incentive provided by Regulation M to maintain Eurodollar holdings. The partial elasticity at the means may be calculated according to the formula:

$$\left(\frac{\partial EDB}{\partial RR^{ed}} \cdot \frac{RR^{ed}}{EDB}\right) = \left(\frac{\delta EDB}{\delta r^{ed}} \cdot \frac{r^{ed}}{EDB}\right)\left(\frac{\delta r^{ed}}{\delta RR^{ed}} \cdot \frac{RR^{ed}}{r^{ed}}\right) + \left(\frac{\delta EDB}{\delta RR^{ed}} \cdot \frac{RR^{ed}}{EDB}\right),$$

the first term is equal to -8.02 x .062 = -.49 which is not offset by the positive second term (.096). Hence a one per cent (one point) rise in the requirement is associated with a .39 per cent ($0.3 billion) decrease in Eurodollar borrowing.

V. Conclusion

Our study confirms that the demand for a particular financial instrument, in our case Eurodollar borrowing, can depend on the returns of a wide range of assets and liabilities available to the transactor. We observe that Eurodollar borrowing is highly sensitive to both the Eurodollar rate itself and several U.S. asset and liability rates. In estimating the responsiveness of Eurodollar borrowing to interest rate variation, we have departed from previous work by incorporating a long lag structure and a wide range of interest rates. The results of our sensitivity analysis indicate these procedures are justified despite a possible presumption of short lags in financial adjustment.

Furthermore our results support the view that these financial instruments are gross substitutes. We also observe that non-price factors have influenced Eurodollar borrowing. Specifically, deviations from trend of bank loans are associated with dramatic changes in the position of these banks in the Eurodollar market.

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TABLE 1

Variable Definitions and Derivations

Dependent Variable

EDB — Liabilities of large city commercial banks to their foreign branches (Monthly average of Wednesday observations, Federal Reserve Bulletin)

Liability Rates

r^{ed} — Three-month Eurodollar deposit rate in London (National Bureau of Economic Research, Data Bank); adjusted for Regulation M Reserve Requirements* (Monthly average of Friday rates)

r^{ff} — Federal funds rate (NBER, Data Bank) (Monthly average of daily rates)

r^{cd} — Rate on 3-month certificates of deposit in New York, end of month (The Economist) adjusted for Regulation Q Reserve Requirements*

Asset Rates

r^{gs} — Rate on 3-5 year U.S. government securities (NBER, Data Bank) (Monthly average of daily rates)

r^{pr} — Prime rate, loans by leading city banks (NBER, Data Bank) (Monthly average of daily rates)

Other Explanatory Variables

RR^{ed} — Regulation M Reserve Requirement on U.S. commercial bank liabilities to foreign branches (Federal Reserve Bulletin)

W — Large city commercial bank (LCB) scale variable (NBER, Data Bank) (Monthly average of Wednesday observations)

$$W = DD(1 - RR^{dd}) + TD(1 - RR^{td}) + \text{Float}$$

DD = demand deposits at LCB's

TD = time deposits at LCB's

RR^{i} = reserve requirement on liability i

L — Gross Adjusted Loans of LCB's (NBER, Data Bank) (Monthly average of Wednesday observations)

We define

$$X_t^* = \sum_{j=0}^{14} (15 - j)X_{t-j}/120 \quad \text{for} \quad X = W, L$$

D1, ..., D11 : Dummies for January, ..., November

Instrumental Variables

Industrial Production Indices for: France, Germany (Federal Republic), Italy, Japan, United Kingdom (IMF, International Financial Statistics)

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*Adjustments to r^{ed} and r^{cd} for reserve requirements followed the formula

$$r^{i}(\text{adjusted}) = r^{i}(\text{unadjusted})/(1 - RR^{i})$$

See Swoboda (1968), for instance.

TABLE 2

Equation 3.1': Estimated Coefficients
Explained Variable: EDB/W*

Variable	Coefficient	t-Ratio
C	0.11734	(13.34)
$1/r^{ed}$	0.07085	(2.92)
$X1^{ed}$	0.00583	(1.23)
$X2^{ed}$	-0.00074	(2.91)
$X1^{ff}$	0.00608	(2.14)
$X2^{ff}$	-0.00008	(0.56)
$X1^{cd}$	-0.00388	(0.64)
$X2^{cd}$	0.00009	(0.28)
$X1^{gs}$	0.01611	(1.97)
$X2^{gs}$	-0.00064	(1.47)
$X1^{pr}$	-0.03209	(7.09)
$X2^{pr}$	-0.00187	(7.12)
$(W-W^*)/W^*$	0.01503	(0.26)
$(L-L^*)/W^*$	0.19128	(2.62)
RR^{ed}	0.03673	(1.93)
D1	0.00012	(0.11)
D2	0.00189	(1.14)
D3	0.00132	(0.80)
D4	-0.00004	(0.03)
D5	0.00050	(0.27)
D6	0.00024	(0.14)
D7	0.00149	(1.18)
D8	0.00220	(1.19)
D9	0.00043	(0.37)
D10	0.00177	(1.55)
D11	0.00250	(2.05)

R^2 = .9947

SSR = .000039

D.W. Statistic = 2.0106

TABLE 3

Point Estimates of Partial Interest Rate Elasticities
$(\partial EDB/\partial r^i \cdot r^i/EDB)$

	Ordinary Least Squares	Instrumental Variables Estimation
r^{ed}	-8.02	-8.02
r^{ff}	5.69	5.26
r^{cd}	-2.23	-1.45
r^{gs}	3.85	2.79
r^{pr}	4.84	5.07

(Computed at $(\overline{1/r})$, $(\overline{EDB/W^*})$)

TABLE 4

Point Estimates of the Own-Rate Elasticity When the Domestic Interest Rate Array is Shortened

Full Array	-8.02
$-r^{ff}$	-7.85
$-r^{cd}$	-8.02
$-r^{gs}$	-7.99
$-r^{pr}$	-6.61
$-(r^{ff}, r^{cd})$	-2.10
$-(r^{gs}, r^{pr})$	-6.11

TABLE 5

Point Estimates of Interest Rate Elasticities and Slopes

	$(\partial EDB/\partial r)(r/E)$	$\frac{\partial EDB/\partial r}{\$B}$
r^{ed}	-8.02	(-5.25)
r^{ff}	5.69	(4.93)
r^{cd}	-2.23	(-1.66)
r^{gs}	3.85	(2.96)
r^{pr}	4.84	(3.47)
Sum of U.S. Liability Rate Elasticities (Coefficients)		
	3.46	(3.27)
Sum of U.S. Asset Rate Elasticities (Coefficients)		
	8.69	(6.43)
Sum of all U.S. Rate Elasticities (Coefficients)		
	12.15	(9.70)